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## ► To cite this version:

Catherine Araujo Bonjean, Jean-François Brun. Chocolate price fluctuations may cause depression: an analysis of price pass-through in the cocoa chain. 2014. halshs-01074157

**HAL Id: halshs-01074157**

**<https://shs.hal.science/halshs-01074157>**

Preprint submitted on 13 Oct 2014

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# **Chocolate price fluctuations may cause depression: an analysis of price pass-through in the cocoa chain**

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*Etudes et Documents n° 20*

First version: March 2010

This version: September 2014

To quote this document:

Araujo Bonjean C., and J.-F. Brun (2010). "Chocolate price fluctuations may cause depression: an analysis of price pass-through in the cocoa chain", *Etudes et Documents*, n° 20, CERDI. [http://www.cerdi.org/back.php/production/show/id/1608/type\\_production\\_id/1](http://www.cerdi.org/back.php/production/show/id/1608/type_production_id/1)

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## Abstract

The aim of this paper is to explore the channels of transmission of the fluctuations in the world price of cocoa to the consumer of chocolate bars in France. This case study can be considered as an illustration of a more general pattern of asymmetric vertical price transmission in the commodity-final product chain. Two types of asymmetry are suspected: asymmetry in the transmission of positive and negative shocks that may reflect non-competitive behavior in the chocolate industry and asymmetry in the transmission of small and large shocks that may be due to adjustment costs. These hypotheses are tested using a three-regime error correction model. Results show that adjustment costs in the processing, manufacturing and distribution of the chocolate tablet are important. Moreover, increases in the cocoa price are more fully and rapidly transmitted to consumers than decreases. These findings may be interpreted as the manifestation of the market power of chocolate companies.

**Key words:** chocolate chain, TAR model, asymmetric price transmission

**JEL codes:** F23, Q13, C32

## Acknowledgment

The authors thank the participants to the First Conference on *Economics and Politics of Chocolate*, Leuven, September 2012, for their comments on a previous version.

## **1. Introduction**

A recent medical research has questioned the causal relationship between chocolate consumption and mood. According to Rose et al. (2010), greater chocolate consumption could be the cause of higher depression. Another and may be more direct subject of concern for chocolate consumers relates to the chocolate price. The commodity price spike of 2007/2008 marks the beginning of a new era of increased instability. The recent and highly publicized episodes of price bursts may give consumers the impression that food prices are always getting higher.

More generally speaking, there is a common perception among consumers that retail prices respond faster to an increase in the price of the raw material than to a downfall. This feeling may be sometimes misleading but it may also reflect an economic reality. For instance, Chen et al. (2005) working on the US gasoline market have shown that the gasoline retail price responds faster to a crude oil increase than to a decrease.

A common feature to the chocolate and the oil industries is to be heavily concentrated at the production and processing stage. Three countries (Côte d'Ivoire, Ghana, Indonesia) account for over 70% of world cocoa production while six multinational companies control about 60% of the world chocolate market. Most of them are vertically integrated. They are engaged in cocoa grinding and the production of final goods such as chocolate bar and confectionery. Moreover, the producer price of cocoa fell dramatically over the past three decades compared to the retail price of chocolate products in the main European consuming markets (Healy, 2005; Gilbert, 2006). The decline in the producer share of the retail value is commonly interpreted as reflecting the monopoly and/or monopsony power of the chocolate industry (Dorin, 2003). Studies conducted in Côte d'Ivoire after the liberalization process have highlighted the lack of competition between traders and processors at the farm-gate level and the low power of small farmers to negotiate prices (Oxfam, 2002).

However, few studies aim at testing the existence of noncompetitive behavior in the cocoa industry and the origin of the producer price decline is still controversial. Among these studies Araujo-Bonjean and Brun (2007) developed a game theory model in which the main cocoa producer, the Côte d'Ivoire, and the chocolate manufacturers compete for price leadership. Econometric estimates of a system of price equations evidence a reversal in price leadership that took place in the mid-80s. The authors conclude that concentration in the processing and manufacturing industry during the 80s has resulted in transferring the price leadership from Côte d'Ivoire to the large chocolate companies. On the contrary, Gilbert

(2006) analyzing the Global Value Chain in the cocoa sector found no evidence of non-competitive behavior in the chocolate industry. He concludes that the fall in the producer share of the retail price of chocolate in the U.K. can be explained by a decrease in production costs in producing countries and an increase in the processing, marketing and distribution costs.

In this paper we reexamine the issue of the impact of market concentration on price formation in the chocolate chain using a new approach focused on the adjustment process of retail prices to cocoa price shocks. The analysis is conducted on an original data set for the French market. A graphical analysis of the relative evolution of the price of cocoa beans and chocolate bar highlights two main phenomena. First, cocoa price fluctuations are passed on to consumers with delays. Secondly, the transmission of price fluctuations appears to be asymmetrical. Sharp upward movements in the cocoa price seem to be more readily transmitted to consumers than price falls. These price adjustment patterns can be taken as evidence of adjustment costs and imperfect competition (Meyer and von Cramon Taubadel, 2004).

An in-depth analysis of the channels of transmission of the fluctuations in the world price of cocoa beans to the consumers of chocolate bars is conducted using threshold cointegration models. Following numerous authors (*e.g.* Balke and Fomby, 1997, Enders and Siklos, 2001, Goodwin and Piggott, 1999), we use the two-step approach to cointegration of Engle and Granger extended to encompass possible asymmetric adjustment to disequilibrium. Two alternatives to the standard linear error correction model are considered. First, short-run price dynamics depends on whether the deviation of the chocolate price from its long run equilibrium is above or below a critical threshold. This asymmetry in the transmission of positive and negative shocks may reflect non-competitive behavior in the chocolate industry. Second, a three-regime threshold autoregressive model catches the existence of adjustment costs. Small shocks are dampened by the processing industry while larger shocks are passed on to consumers. Thus, price transmission is asymmetric according to the sign of the shock and to the size of the shock.

Estimations are conducted on a sample of monthly prices over the period January 1960 to February 2003. Robustness tests are conducted on a sample of annual prices running from 1949 to 2011. Results show that the chocolate price does not adjust to small shocks in the cocoa market but adjusts to large deviations. Moreover, the speed of adjustment is larger for negative deviations than for positive ones.

The paper is organized as follows. Section two presents the evolution of the price of cocoa beans and chocolate bar and briefly highlights the main features of the chocolate marketing chain. Section three exposes the model of vertical price transmission under consideration and the testing strategy. Section four presents the econometric results. The last section concludes.

## 2. The cocoa – chocolate chain: main features

The comparison of commodity and manufactured goods prices may be misleading because of the changing nature of the final goods over time. In the cocoa sector, changing consumer tastes and technological progress may have induced shifts in the composition and cost structure of chocolate products while cocoa bean is a homogenous good. Thanks to an original data set we are able to avoid this potential pitfall. The evolution of the retail price of a homogenous chocolate bar sold on the French market can be traced since 1949. It is a 100 grams bar of black chocolate (“tablette”) with unchanged characteristics<sup>1</sup>. The price of this chocolate bar is available on an annual basis over the period 1949 – 2011 and on a monthly basis over the period January, 1960 to February, 2003.

The relative evolution over 1955 – 2011 of the world price of cocoa beans<sup>2</sup> and the retail price of the chocolate bar in the French market highlights two phenomena (figure 1). First, cocoa price fluctuations are passed on to the price of the chocolate bar with a lag of roughly one year. Second, positive shocks in the cocoa price appear to be fully transmitted to the retail chocolate price. This phenomenon can be especially observed during the 70s. By contrast, cocoa price decreases are passed on to the chocolate price in a lesser way or are not transmitted at all. For instance, the cocoa price experienced a sharp drop during the second half of the 80s which was not fully transmitted to the chocolate price. Moreover, from 1985 until the end of the 90s, the cocoa price experienced a long lasting phase of decline while the chocolate price stayed at a rather steady level.

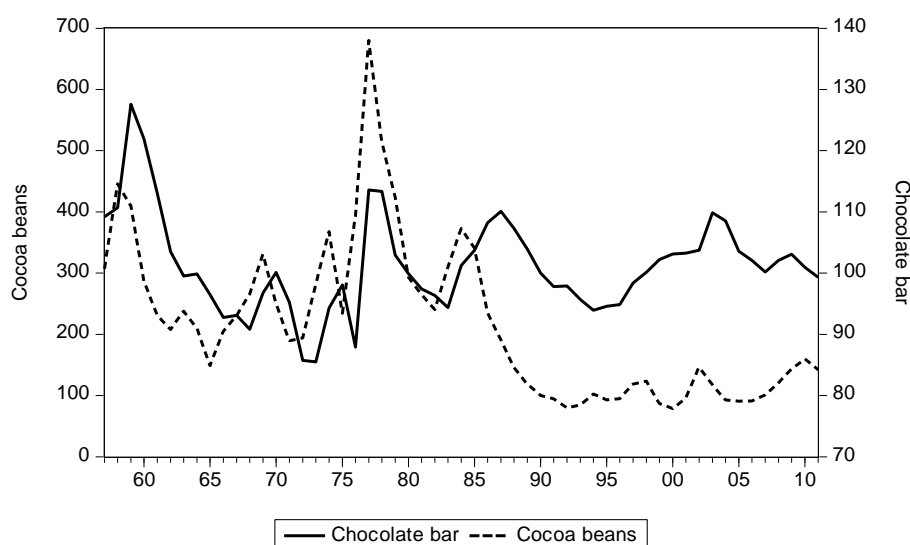
These features of price movements in the cocoa chain suggest the existence of important lags and asymmetries in price transmission. Increases in the input price seem to be more fully transmitted to the output price than equivalent decreases.

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<sup>1</sup> A 100 grams bar of chocolate contains approximately 60 grams of cocoa bean. Source INSEE, France.

<sup>2</sup> Average of London and New York Stock Exchange prices. Source: International Financial Statistics of the IMF.

Figure 1. Evolution of the price of cocoa beans and chocolate bar over 1957 – 2011



Prices are expressed in constant euros (100 = 1990).

Source: INSEE and IMF

Among the possible causes for asymmetric price transmission, two main factors have received a special attention in the literature: a non-competitive market structure and adjustment costs. These two factors may result in two different types of asymmetry in the adjustment process.

#### *Asymmetry in the transmission of positive and negative shocks*

This type of asymmetry may be due to imperfect competition in the processing/distribution chain. The agro-food industry is highly concentrated and processing companies and/or distributors are often accused to exert an excessive market power. Processing industry may be prompted to pass on increases in input prices to output prices more rapidly and more fully than decreases.

The cocoa industry has experienced important changes during the period under consideration, leading to a significant increase in industry concentration. Every stage of the cocoa chain - from cocoa bean production to chocolate products distribution - is highly concentrated. The three largest cocoa producing countries (Côte d'Ivoire, Ghana, Indonesia) account for more than 70 % of cocoa world production while three trading and grinding companies and six manufacturers dominate the chocolate processing, manufacturing and distribution.



At the world level three companies control about 40 % of cocoa grinding<sup>3</sup>: Cargill (14%), Archer Daniels Midland (ADM) (14%) and Barry Callebaut (12%). Petra Foods and Blommer play also an increasing role, with respectively 7% and 5% of world grinding capacities<sup>4</sup>. Except Cargill and ADM whose core activity is trade and grinding, the other companies are engaged in the production of industrial chocolate<sup>5</sup> and semi-finished products. Barry Callebaut, the largest producer of industrial chocolate (about 40% of the world market for industrial chocolate) is also moving into the manufacture of consumer chocolate. Cocoa grinding generate low added value. Cocoa butter and power are traded on world exchanges at a price which is tightly linked to the price of the cocoa bean.

Six manufacturers dominate the world market of chocolate and confectionery products in 2008: Mars (15%), Nestlé (12.5%), Kraft Foods (8.1%), Cadbury Schweppes (7.2%), Hershey (7.2%) and Ferrero (7.2). With Lindt (3.6%) they are estimated to control more than 60% of the market for finished chocolate products<sup>6</sup>. These groups which are vertically integrated produce the industrial chocolate they need to make consumer products. The merger and acquisition process is still expanding with the acquisition of Cadbury by Kraft Foods in 2010.

In France, the market segment for industrial chocolate is dominated by Barry Callebaut and Cémoi (Cantalou) while five major manufacturers controlling about 70 % of the consumer chocolate market: Ferrero (22%), Kraft Foods (15%), Lindt (14%), Nestlé (14%) and Mars (10%)<sup>7</sup>. The distribution sector is also highly concentrated. Hypermarkets and supermarkets are the main dealers of chocolate products. They sell the quasi totality (80%)<sup>8</sup> of the chocolate bars consumed in France. This sector also experienced an increased concentration during the last two decades.

#### *Asymmetry in the transmission of large and small shocks*

Adjustment costs in the packaging and distribution stage of the marketing process are a possible cause of asymmetry in price transmission according to the size of the shocks. Changing prices generate so-called “menu costs” (Barro, 1972). For instance the costs of reprinting price lists or catalogues may lead to late and asymmetric adjustment of prices. Fixed adjustment costs are expected to create a price band inside which the retail price does

<sup>3</sup> During this process cocoa beans are transformed successively into nibs, liquor then butter and powder. Liquor and butter are semi-finished products used as ingredients in chocolate bars.

<sup>4</sup> Source: Oxfam (2008) and UNCTAD (2008). Data refer to the 2006/2007 campaign.

<sup>5</sup> Industrial chocolate is the raw material of consumer chocolate manufacturers.

<sup>6</sup> Source: Enjeux Les Echos n° 271, 01 Septembre 2010.

<sup>7</sup> Data refer to the 2009/10 campaign. Source :Lineaires, Octobre 2010.

<sup>8</sup> Source : UNCTAD (2008)

not adjust to fluctuations in the raw material price as the adjustment cost would exceed the benefit. As a consequence, processors and/or distributors respond to “small” input price fluctuations by increasing or reducing their margins. The output price will adjust only if the fluctuations in the input price exceed a critical level. Moreover, the adjustment costs are not necessarily symmetric with respect to price increases or decreases (see for instance: Meyer and von Cramon-Taubael, 2004, Peltzman, 2000).

Non-competitive markets and adjustment costs may result in a nonlinear price dynamic. On the one side, threshold effects occur when large shocks bring about a different response than small shocks due for instance to the presence of adjustment cost. On the other side, threshold effects occur when positive shocks (increases in the input price) trigger a different response of the output price than negative shocks. In the next section we develop a nonlinear model of price transmission that takes into account these two types of asymmetry.

### 3. Model of vertical price transmission and estimation procedure

As the price series of the cocoa beans and the chocolate bar are I(1) (see table 1 below), we examine the relationship between the two series in a cointegration framework. While standard models of cointegration assume linear and symmetric adjustment towards the long run equilibrium relationship, we will consider a nonlinear model of cointegration with asymmetric adjustment.

#### *Standard linear and symmetric adjustment model*

The long run relationship between the two prices is given by:

$$Pc_t = \alpha_0 + \alpha_1 Pb_t + \mu_t \quad (1)$$

$Pc_t$  is the price of the chocolate bar at time  $t$ .  $Pb_t$  is the price of the cocoa beans.

Cointegration of  $Pc$  and  $Pb$  depends upon the nature of the autoregressive process for  $\mu_t$ . In the standard model of cointegration  $\mu_t$  is stationary with zero mean and may follow a linear AR(p) process in the form:

$$\mu_t = \sum_{i=1}^p \gamma_i \mu_{t-i} + e_t \quad (2)$$

$e_t$  is a white noise disturbance.

If the series are cointegrated, the Granger representation theorem guarantees the existence of an error-correction representation of the variables in the form:

$$\Delta P_t = \lambda \mu_{t-1} + \sum_{j=1}^p \theta_j \Delta P_{t-j} + \varsigma_t \quad (3)$$

with  $P_t = (P_{Ct}, P_{Bt})'$  and  $\zeta_t$  a white noise disturbance.  $\mu_{t-1}$  is the error correction term.  $\lambda = (\lambda_{Pc}, \lambda_{Pb})$  is the vector of the adjustment speed of  $\Delta P_t$  to a deviation of  $P_t$  from its long-run equilibrium level.

In this standard error correction model (ECM), adjustment is linear and symmetric:  $\lambda_{pc}$  and  $\lambda_{pb}$  are constant, with  $|1 + (\lambda_{pc} - \alpha_1 \lambda_{pb})| < 1$ . At every period, a constant proportion of any deviation from the long run equilibrium is corrected, regardless of the size or the sign of the deviation and the system moves back towards the equilibrium.

However, as discussed above, different types of market failures may prevent a continuous and linear adjustment of prices. Threshold cointegration model are thus considered.

*Nonlinear and asymmetric adjustment: the threshold autoregressive model*

If  $Pc$  and  $Pb$  are characterised by asymmetric adjustment, the equilibrium error,  $\mu_t$ , can be modeled as a self-exciting threshold autoregressive (TAR) process. In such a model the autoregressive decay depends on the state of the variable of interest,  $\mu_{t-d}$ .

The general model for the equilibrium error, allowing for nonzero intercept and asymmetric thresholds is given by a TAR(k; p, d) model of the form:

$$\mu_t = \gamma_0^{(j)} + \sum_{i=1}^p \gamma_i^{(j)} \mu_{t-i} + e_t^{(j)}, \quad r_{j-1} \leq \mu_{t-d} < r_j \quad (4)$$

where  $k$  is the number of regimes that are separated by  $k-1$  thresholds  $r_j$  ( $j = 1$  to  $k-1$ ). In each regime,  $\mu_t$  follows a different linear autoregressive process depending on the value of  $\mu_{t-d}$ .  $d$  is the threshold lag or delay parameter. It represents the delay in the error correction process when agents react to deviations from the equilibrium with a lag.  $p$  is the order of the autoregressive process.  $e_t^{(j)}$  are zero-mean, constant-variance *i.i.d* random variables.

The stationarity of the threshold autoregressive process depends on the behavior of  $\mu_t$  in the outer regimes. The equilibrium error may behave like a random walk inside the threshold range, but as long as it is mean-reverting in the outer regimes it is a stationary stochastic process (Balke and Fomby, 1997).

We focus on a three regime model with two thresholds allowing asymmetric adjustment to deviations in the positive and negative directions taking into account asymmetry in the correction of large and small deviations. This pattern of adjustment takes the form of a TAR(3; p, d) model:

$$\mu_t = \begin{cases} \gamma_0^l + \sum_{k=1}^p \gamma_k^l \mu_{t-k} + v_t^l & \text{if } \mu_{t-d} < \tau_1 \\ \gamma_0^m + \sum_{k=1}^p \gamma_k^m \mu_{t-k} + v_t^m & \text{if } \tau_1 \leq \mu_{t-d} \leq \tau_2 \\ \gamma_0^u + \sum_{k=1}^p \gamma_k^u \mu_{t-k} + v_t^u & \text{if } \mu_{t-d} > \tau_2 \end{cases} \quad (5)$$

$v_t$  are zero-mean random disturbances with constant standard deviation.  $\tau_1$  and  $\tau_2$  are the unknown threshold values.

The corresponding vector error correction representation is given by:

$$\Delta P_t = \begin{cases} \pi_0^l + \phi^l \mu_{t-1} + \sum_{k=1}^p \gamma_k^l \Delta P_{t-k} + \zeta_t^l & \text{if } \mu_{t-d} < \tau_1 \\ \pi_0^m + \phi^m \mu_{t-1} + \sum_{k=1}^p \gamma_k^m \Delta P_{t-k} + \zeta_t^m & \text{if } \tau_1 \leq \mu_{t-d} \leq \tau_2 \\ \pi_0^u + \phi^u \mu_{t-1} + \sum_{k=1}^p \gamma_k^u \Delta P_{t-k} + \zeta_t^u & \text{if } \mu_{t-d} > \tau_2 \end{cases} \quad (6)$$

The two thresholds  $\tau_1$  and  $\tau_2$  define three price regimes. The matrix of parameters  $\phi^l$ ,  $\phi^m$ , and  $\phi^u$ , give the adjustment speeds of one price to deviations from the equilibrium relationship. The speed of adjustment differs according to whether the deviation from long run equilibrium is above or below the critical thresholds.

A case often encountered is when  $\phi^m = 0$ . In that case, small deviations from equilibrium are not corrected. Deviations from the equilibrium must reach a critical level before triggering a price response.

If  $|\tau_1| \neq |\tau_2|$  the interval  $[\tau_1, \tau_2]$  is not symmetric around the origin and deviations in the positive and negative directions must reach different magnitudes before triggering a price response. This case is more likely when adjustment costs are asymmetric.

### *Testing strategy*

To characterize the relationship between the cocoa price and the chocolate price we follow the threshold cointegration method introduced by Balke and Fomby (1997) and developed by Enders and Siklos (2001). It is a two-step approach that extends the Engle and Granger (1987) testing strategy by permitting asymmetry in the adjustment towards equilibrium. The first step involves the estimation of the long run equilibrium relationship between the price of chocolate and the price of cocoa; cointegration tests are applied to the equilibrium error. The second step involves estimating the best fit TAR model for the

equilibrium error and testing for nonlinear threshold behaviour. Tests for significant difference in parameters across alternative regimes are conducted and confidence intervals for the threshold values are calculated. Lastly, the asymmetric error correction model corresponding to the best fit TAR model is estimated.

The empirical analysis is conducted successively on monthly and annual data. The first sample includes high frequency prices that are only available over the period January 1960 to February 2003 (518 observations). The second sample covers a longer period of time, 1949 – 2011, but with a lower number of observations (63). This annual sample is used to conduct some robustness tests.

## 4. Econometric results

### 4.1. Testing for no cointegration against linear cointegration

We first test for the order of integration of price series using Dickey and Fuller (1981), Phillips and Perron (1988), and Kwiatkowski et al. (1992) tests (table 1). The ADF and PP tests fail to reject the null hypothesis of unit root for both the cocoa and the chocolate bar price series, in annual and monthly frequency, while the first-differenced price series appears to be stationary<sup>9</sup> (table 1).

Table 1. Unit root tests.

	ADF <sup>a</sup>	Prob	PP	Prob	KPSS
<i>Monthly data, 1960.01 – 2003.02</i>					
Price of the cocoa bean ( <i>Pb</i> )	-1.88	0.34	-1.84	0.36	0.44***
Price of the chocolate bar ( <i>Pc</i> )	-2.20	0.49	-2.16	0.51	0.33***
<i>Annual data, 1949 – 2011</i>					
Price of the cocoa bean ( <i>Pb</i> )	-2.998	0.141	-2.383	0.385	0.125*
Price of the chocolate bar ( <i>Pc</i> )	-2.546	0.306	-1.895	0.645	0.166**
Consumer Price Index ( <i>CPI</i> )	-2.816	0.198	-2.065	0.553	0.117

ADF: Augmented Dickey Fuller test. Ho: unit root

PP: Phillips- Perron. Ho: unit root

KPSS: Ho: I(0) ; critical values (Kwiatkowski-Phillips-Schmidt-Shin, 1992, Table 1).

<sup>a</sup>Lag order selection based on the Akaike Information Criterion. Trend and intercept in test equation.

\*: significant at the 10% level; \*\*: significant at the 5% level; \*\*\*: significant at the 1% level.

<sup>9</sup> Results not reported here.

The long run relationship between the chocolate and the cocoa price is estimated using Fully Modified OLS (Phillips and Hansen, 1992)<sup>10</sup>; the estimated equation is given in table 2. The Johansen procedure is also conducted to test for linear cointegration between the two price series as well as the exogeneity of the price of cocoa beans (table 3). Indeed, the cocoa and the chocolate price may be endogenously determined if the concentration in the chocolate industry allows large manufactures to exert a monopoly and monopsony power.

Table 2. FMOLS estimate of the cointegrating equation. Dependent variable:  $P_c$

intercept	$P_b$	Trend	Adj R <sup>2</sup>	No obs	Engle-Granger cointegration tests	
					t-stat	z-stat
-0.129	0.219	0.002	0.947	517	-2.133	-5.810
(0.000)	(0.098)	(0.000)			(0.713)	(0.895)

Monthly data on the sample: 1960.01– 2003.02

P-value are in parentheses. Number of lags in the Engle-Granger test equation: 3

z is the normalized autocorrelation coefficient for residuals

Table 3. Johansen cointegration tests.

Trace test	Max eigenvalue	Max eigenvalue and trace test
r=0	r=0	r=1
35.61	26.18	9.42
(0.002)	(0.004)	(0.16)

Monthly data, 1960.01 – 2003.02

H<sub>0</sub>: no cointegration, intercept and trend in the cointegrating equation, no intercept in VAR

Results of the tests of no cointegration against linear cointegration are contrasted: the residual-based Engle-Granger test fails to reject the null hypothesis of no cointegration between the cocoa and the chocolate price series (table 2) while the Johansen tests clearly reject the null of no cointegration (table 3).

<sup>10</sup> Fully Modified OLS (FMOLS) account for serial correlation and for the endogeneity in the regressors that results from the existence of a cointegrating relationship.

Table 4. Exogeneity test results from VECM

Excluded in		$\Delta P_c$	$\Delta P_b$
Error correction term	chi-sq	15.297	1.605
	(p-value)	(0.0001)	(0.205)
Lagged $\Delta P_b$	chi-sq	16.54	
	(p-value)	(0.002)	
Lagged $\Delta P_c$	chi-sq		2.819
	(p-value)		(0.589)

Intercept and trend in the cointegrating equation, no trend in VAR.

Number of lags selected according to Schwarz criterion. Lags interval in first difference: 1 to 4

The tests on the adjustment coefficients in the VECM representation do not reject weak exogeneity of the cocoa price with respect to the cointegrating vector parameters (table 4). Moreover the block exogeneity Wald (or Granger causality) tests do not reject strong exogeneity of the cocoa price. According to these results the retail price of chocolate adjusts to past cocoa price while the reverse is not true. These findings do not support the monopsony power hypothesis. The chocolate manufacturers do not seem to be able to exert pressure on the world price of cocoa beans. In what follows, the cocoa price is considered to be exogenous and the error correction model is restricted to the chocolate bar price equation.

#### 4.2. Testing for no cointegration against threshold cointegration

With a  $t$ -stat equal to -2.13 the Engel-Granger test (table 5) would lead to the non-rejection of the null of no cointegration at the conventional level. However, cointegration test may have low power if data are generated by a TAR model<sup>11</sup>. We thus estimate a univariate TAR model with one threshold for the residuals of the cointegrating equation and test the null of no cointegration against the alternative of a stationary TAR model.

The test equation corresponding to a TAR model with one threshold is given by:

$$\Delta \mu_t = I_t \rho_1 \mu_{t-1} + (1 - I_t) \rho_2 \mu_{t-1} + \sum_{i=1}^{p-1} \gamma_i \Delta \mu_{t-i} + \varepsilon_t \quad (7)$$

$I_t$  is an indicator function that depends on the level of  $\mu_{t-d}$ .  $I_t = 1$  if  $\mu_{t-d} \geq \tau$  and  $I_t = 0$  otherwise.  $\tau$ : is the threshold value.  $\varepsilon_t$  is an *iid* process with zero mean and constant variance.

<sup>11</sup> According to Enders and Granger (1998) and Enders and Siklos (2001) the power of the Engle-Granger test generally exceeds that of the  $\Phi$  and  $t$ -Max statistics when the true data-generating process has only one single threshold. However, the power of the  $\Phi$  stat may be higher than standard unit root tests in two threshold models.

The test statistics for threshold cointegration are the  $t_{\text{MAX}}$  statistic given by the larger  $t$ -statistic of  $\rho_1$  and  $\rho_2$ , and the  $F$ -statistic, called  $\Phi$  stat, corresponding to the null hypothesis  $\rho_1 = \rho_2 = 0$ . The threshold value is first set equal to zero and the delay parameter to one; in a second stage these parameters are estimated along with the value of  $\rho_1$  and  $\rho_2$ . Results for the tests are given in table 5.

Table 5. Nonlinear cointegration test results. Dependent variable:  $\mu_t$ 

	Engle-Granger test equation	TAR(2; 4,1) <sup>a</sup> $\tau = 0$	TAR(2; 4, 8) <sup>b</sup> Consistent threshold and delay parameter.
$\rho_1$	-0.005 (-2.130)	-0.004 (-1.249)	-0.030 (-4.939)
$\rho_2$	-	-0.005 (-1.732)	-0.001 (-0.328)
$\gamma_1$	0.392 (8.906)	0.392 (8.880)	0.365 (8.348)
$\gamma_2$	0.085 (1.792)	0.084 (1.779)	0.079 (1.706)
$\gamma_3$	0.112 (2.540)	0.111 (2.514)	0.105 (2.426)
Schwarz criterion	-8.982	-8.970	-9.001
$\Phi$	-	2.299	12.264
Threshold value		0	-0.089
$\rho_1 = \rho_2$			20.130
[p-value]			[0.000]
$F_{12}$			20.972
[p-value]			(0.042) <sup>c</sup>
F heteroskedasticity test	9.476		10.433
[p-value]	[0.000]		[0.000]

$t$  statistics are in parenthesis.

a. TAR process with 2 regimes, 4 lags and delay parameter equal to 1.

Critical value for  $t_{\text{MAX}}$  (Enders and Siklos, 2001, table 2): 90 %: -1.69; 95 %: -1.89; 99 %: -2.29

Critical value for  $\Phi$  (Enders and Siklos, 2001, table 5): 90 %: 5.21; 95 %: 6.33; 99 %: 9.09

b. TAR process with 2 regimes, 4 lags and delay parameter equal to 8.

Critical value for  $t_{\text{MAX}}$  (Enders and Siklos, 2001, table 6): 90 %: -1.52; 95 %: -1.73; 99 %: -2.30

Critical value for  $\Phi$  (Enders and Siklos, 2001, table 5): 90 %: 6.44; 95 %: 7.56; 99 %: 10.16

<sup>c</sup> Bootstrapped p-value for 500 replications.



Assuming a threshold value equal to zero and a delay parameter equal to one - TAR(2; 4,1) model, column 3, table 5 - the estimated speed of adjustment  $\rho_1$  and  $\rho_2$  are negative but the  $\Phi$  stat (equal to 2.299) is less than the critical value while the  $t_{MAX}$  (-1.249) is larger than the critical value at the conventional level. We thus cannot reject the null of no cointegration for  $\tau = 0$ .

To find the consistent estimate of the threshold ( $\tau$ ) along with the delay parameter ( $d$ ) we use Chan's (1993) method. It consists in minimizing the sum of squared residuals by searching over a set of potential threshold values for all possible delay parameters. The lagged residuals of the long run relationship are sorted in ascending order and the largest and smallest 7.5 % of the residuals are discarded. The remaining 85 % of the residual values are considered as potential thresholds. The threshold value and the delay parameter yielding the lowest sum of squared residuals are considered as the appropriate estimates of the threshold value and delay parameter.

The estimation results are presented in the last column of table 5. The estimated delay parameter and threshold value corresponding to the lowest sum of squared error are respectively equal to 8 and -0.089.  $\rho_1$  and  $\rho_2$  are now jointly significantly different from zero at the 1% level. The speed of adjustment significantly differs according to the regime. It is large when the disequilibrium term falls below the threshold and close to zero in the other regime. We thus reject the null of no cointegration of the price series against an asymmetric adjustment process.

#### 4.3. Testing nonlinearity

To test for threshold nonlinearity in the cointegrating residual  $\mu_t$ , we implement the sup- $F$  test developed by Hansen (1997, 1999). The testing procedure is based on nested hypothesis tests. It consists to test the null of linearity, or TAR(1), against the alternative of TAR( $m$ ) model using a LR-type test of the form:

$$F_{1m} = n \left( \frac{S_1 - S_m}{S_m} \right)$$

$S_1$  is the sum of squared residuals under the null of linearity.

$S_m$  is the sum of squared residuals under the alternative hypothesis of a  $m$ -regime TAR( $m$ ) model with  $m > 1$ .

Testing linearity versus a threshold alternative involves a non-standard inference problem as the threshold parameters are not identified under the null hypothesis. In such case

conventional test statistics do not have standard distribution. We use Hansen (1996) bootstrap procedure to approximate the asymptotic distribution of  $F$ .<sup>12</sup> Because there is evidence of heteroskedasticity in the residuals of the TAR(1) and TAR(2) models (White heteroskedasticity test, table 5) we replace the  $F$ -statistic  $F(\tau)$  with a heteroskedasticity-consistent Wald statistic and modify accordingly the bootstrap procedure (Hansen, 1997).

TAR(1) model is tested against TAR(2) and TAR(3). As previously described, the three regimes TAR model is fitted to  $\mu_t$  by minimizing the sum of squared residuals with respect to the threshold and delay parameters, maintaining the lag length at four. Following Goodwin and Piggott (1999), a two dimensional grid search is used, to estimate the two thresholds  $\tau_1$  and  $\tau_2$ . As a practical matter, we search for the first (second) threshold within negative (positive) residuals dropping the 5% largest and smallest ones.

$F_{12}$  is equal to 20.97 (table 5) and  $F_{13}$  to 32.05 (table 6) with a bootstrapped p-value respectively equal to 0.042 and 0.028 leading to the rejection of the TAR(1) model.

#### 4.4. The selected threshold model

We now turn to the estimation of the best fit TAR( $m$ ) model. To determine the appropriate number of regimes,  $m$ , we use the  $F_{23}$  statistic proposed by Hansen (1999):

$$F_{23} = n \left( \frac{S_2 - S_3}{S_3} \right)$$

$S_2$  is the sum of squared residuals under the null of TAR(2) model.

$S_3$  is the sum of squared residuals under the alternative hypothesis of a three-regime TAR(3) model.

Table 6. Estimate results for the 3 regimes - TAR(3; 4, 8) - model. Dependent variable:  $\Delta\mu_t$

	regime			Threshold value	Asymmetry tests		Linearity tests	
	lower	middle	upper	[confidence interval]	$\rho^l = \rho^u$	$\rho^l = \rho^m = \rho^u$	$F_{13}$	$F_{23}$
$\rho$	-0.031	0.006	-0.009	-0.089	4.119	10.609	32.048	10.644
(t-stat)	(-2.975)	(2.025)	(-3.733)	[-0.0896 ; -0.088] <sup>a</sup>	(0.043)	(0.000)	(0.028) <sup>c</sup>	(0.086) <sup>c</sup>
No obs	[27]	[403]	[79]	0.066				
				[0.06217 ; 0.07598] <sup>b</sup>				

TAR model with 3 regimes, four lags and a delay parameter equal to 8.

White Heteroskedasticity-Consistent Standard Errors & Covariance

$t$  statistics in parenthesis.

<sup>a</sup> 95 % asymptotic confidence interval for  $\tau_1$ .

<sup>12</sup> 500 simulations are run replacing the dependant variable by standard normal random draws. For each bootstrapped series the grid search procedure is used to estimate the threshold value and the  $F_n(\tau)$  statistic is computed.

<sup>b</sup> 90 % asymptotic confidence interval for  $\tau_2$ .

<sup>c</sup>: Bootstrapped-calculated asymptotic P-value using 500 replications.

We also use the Hansen (1996) methodology to construct confidence-intervals for the threshold estimates.

$$LR_n(\tau) = n \left( \frac{\tilde{\sigma}_n^2(\tau) - \hat{\sigma}_n^2(\hat{\tau})}{\hat{\sigma}_n^2(\hat{\tau})} \right)$$

$\tilde{\sigma}^2(\tau)$  is the residual variance given the true value of the threshold.

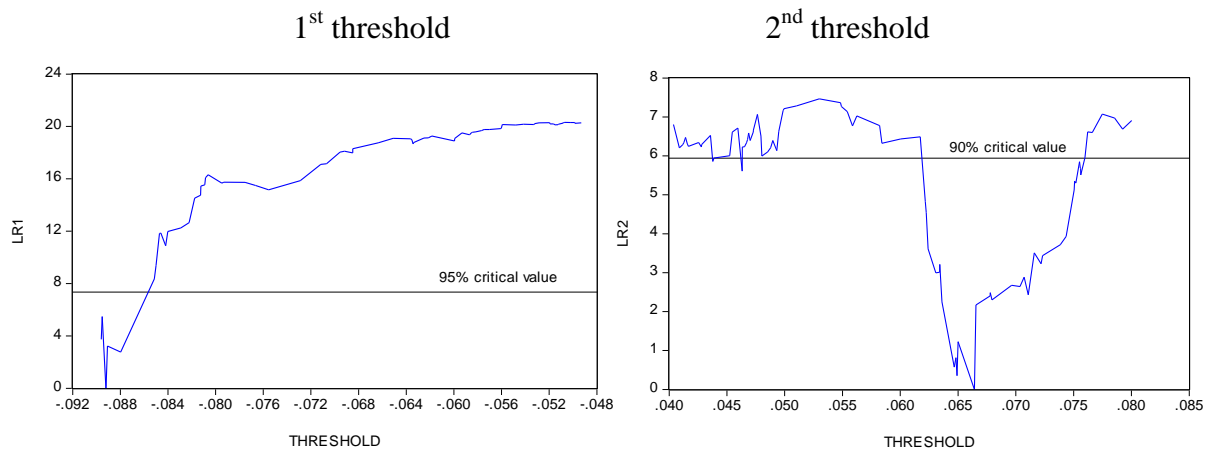
$\hat{\sigma}^2(\tau)$  is the residual variance given the estimated value of the threshold.

$LR_n(\tau)$  is the likelihood ratio statistic to test the hypothesis  $\tau = \tau_0$ . The  $\beta$ -percent confidence interval for  $\tau$ , is given by:  $\hat{\Gamma} = \{\tau : LR_n(\tau) \leq c(\beta)\}$  where  $c(\beta)$  is the  $\beta$ -level critical value from the asymptotic distribution of  $LR_n(\tau)$ . Following Hansen (1997) we use the convexified region  $\hat{\Gamma}^c = [\hat{\tau}_1, \hat{\tau}_2]$  where  $\hat{\tau}_1$  and  $\hat{\tau}_2$  are the minimum and the maximum elements of  $\hat{\Gamma}$ .

In the heteroskedastic case, the modified likelihood ratio is:  $LR_n^*(\tau) = n \left( \frac{\tilde{\sigma}^2(\tau) - \hat{\sigma}^2}{\eta^2} \right)$

and the modified likelihood ratio confidence region is  $\hat{\Gamma} = \{\tau : LR_n^*(\tau) \leq c(\beta)\}$ . The nuisance parameter  $\eta$  is estimated using a polynomial regression (Hansen, 1997).

Figure 2. Confidence interval for the estimated threshold



The value for  $F_{23}$  is 10.644 (table 6) leading to the rejection of the null of a two-regime TAR model against the alternative of a three-regime TAR model although at the 10% confident level only.

The  $F$ -test for equality of the adjustment parameters across regimes rejects the hypothesis of symmetric adjustment ( $\rho^l = \rho^u$  and  $\rho^l = \rho^m = \rho^u$  in table 6). The first regime includes 33 observations, the second 400 and the third 75 observations.

As long as  $\mu_{t-8}$  is inside the band - defined by the two thresholds -  $\mu_t$  acts as a unit root process and consequently has no tendency to drift back towards some equilibrium. When  $\mu_{t-8}$  is above  $\tau_2$ ,  $\mu_t$  becomes an  $I(0)$  process which tends to revert back to the upper border of the band. In the same way, when  $\mu_{t-8}$  is below  $\tau_1$ ,  $\mu_t$  is  $I(0)$  and tends to revert even quicker to the lower border of the band.

The estimated threshold values are -0.086 and 0.069. A plot of the adjusted likelihood ratio ( $LR_n^*(\tau)$ ) for the two thresholds is displayed in figure 2. This figure shows that the first threshold estimate is quite precise while the confidence interval for the second threshold is much larger. The two thresholds are not symmetric around zero suggesting that negative deviations from the long run equilibrium must reach a higher level (in absolute value) than positive deviations before triggering a response in the chocolate price.

#### 4.5. Short run dynamic

The error correction model shows that the chocolate price displays asymmetric error correction toward long run equilibrium (table 7). The price of chocolate adjusts faster to negative shocks than to positive shocks. In the middle regime the coefficient for the error correction term is not significantly different from zero. Deviations of the chocolate price from its long run equilibrium thus have to reach a critical level before adjustment operates. For small deviations the chocolate and the cocoa prices move independently.

Table 7. Estimate results for the 3 regimes Error Correction Model. Dependent variable:  $\Delta P_c$

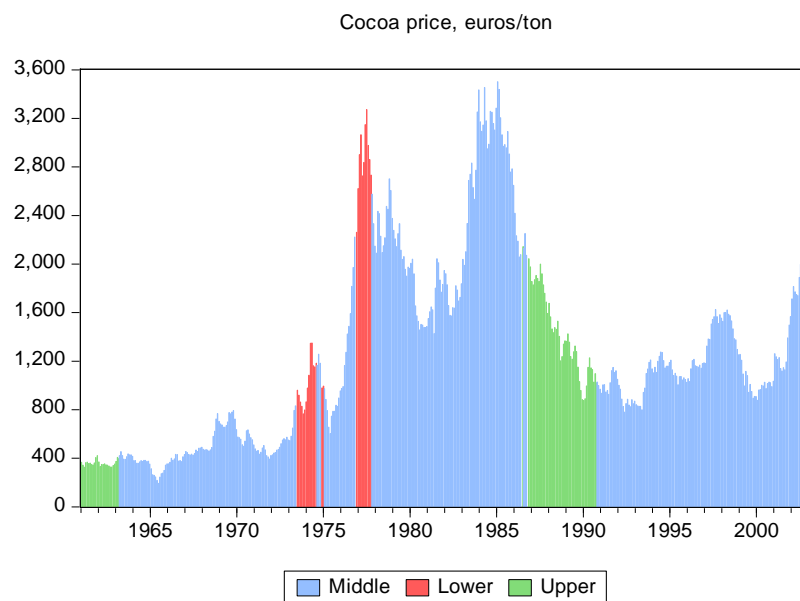
regime						
lower	middle	upper	Adj R <sup>2</sup>	Wald tests ( $F$ -stat)		
$\phi^l$	$\phi^m$	$\phi^u$		$\phi^l = \phi^m = \phi^u = 0$	$\phi^l = \phi^m = \phi^u$	$\phi^l = \phi^u$
-0.029	0.003	-0.008	0.369	6.669	7.472	5.012
(0.001)	(0.339)	(0.004)		(0.000)	(0.001)	(0.026)

White Heteroskedasticity-Consistent Standard Errors & Covariance. 3 lags in first difference variables.  
P-value in parenthesis.

Figure 3 depicts the regime shifts over the period under consideration. During the major part of the period, the deviations from the long run equilibrium relationship - linking the price of chocolate to the price of cocoa - fell inside the band regime (blue period). Within

the band prices are not cointegrated. During two short periods of time, 1973-1975 and 1977-1979, corresponding to two successive booms in the world cocoa prices, the deviation from the long run equilibrium fell below the first threshold. The chocolate bar price was well below its long run equilibrium value and tended to revert back rapidly whereas, the 1987-1991 period corresponded to a phase of low cocoa prices. The chocolate price was above its long run equilibrium value and tended to move back toward the equilibrium but rather slowly.

Figure 3. Timing of regime switching



#### 4.6. Robustness tests

To test the robustness of the results the analysis is duplicated on a sample of annual prices covering the 1949 – 2011 period. Two specifications of the long run relationship between the chocolate bar and the cocoa prices are considered. The first one is the same as the one tested on monthly prices *i.e.* a linear model with two variables: the price of cocoa beans and the price of the chocolate bar. The second specification is a log linear model including an additional variable - the consumer price index in France – that catches the evolution of the cost of inputs entering in the chocolate making process.

##### *Testing no cointegration against linear cointegration*

The FMOLS estimate of the long run relationship is given in table 9 as well as cointegration test results. Both specification evidence a long run positive relationship between

the cocoa and the chocolate bar prices with a long run elasticity of the chocolate price to cocoa price equal to 7.5 % (model 2 table 8).

The Engle-Granger (EG) cointegration tests reject the no cointegration hypothesis for the bivariate linear model (model 1) as well as for the multivariate log linear model (model 2). These results are supported by the Johansen tests which also clearly reject the null of no cointegration.

Table 8. FMOLS estimate of the cointegrating equation and cointegration tests.

Annual data, 1949 – 2011. Dependent variable:  $P_c$

Model specification	$P_b$	CPI	intercept	$R^2$	EG cointegration test			Johansen coint tests		
					No obs	t-stat	z-stat	Trace	max eig.	Max eig and trace
Model 1	0.046 (0.001)		yes	0.306	62	-1.113 (0.013)	-2.773 (0.001)	r = 0 21.107 (0.006)	r = 0 21.078 (0.004)	r = 1 0.029 (0.865)
Model 2	0.075 (0.000)	0.939 (0.000)	no	0.984	56	-5.918 (0.000)	-23.892 (0.043)	r = 0 36.880 (0.007)	r = 0 26.527 (0.008)	
								r = 1 10.353 (0.255)	r = 1 8.630 (0.318)	r = 2 1.723 (0.189)

$P_c$ : chocolate bar price;  $P_b$ : cocoa bean price; CPI: consumer price index in France.

Johansen trace and max eigenvalue tests:  $H_0$ : no cointegration, linear trend in data, intercept (no trend) in the cointegrating equation and VAR.

$r$  = hypothesized number of cointegrating equation. Two lags in VECM.

### Testing nonlinearity

The no cointegration hypothesis being rejected by the Engle-Granger test we directly turn to the nonlinearity tests. Test results are given in table 9.

Table 9: Non-linearity tests results

model	Y	X	F12	F13	F23
Model 1	$P_c$	$P_w$	5.484 (0.080)	18.376 (0.014)	11.844 (0.065)
Model 2	$\log(P_c)$	$\log(P_w), \log(IPC)$	4.014 (0.192)	11.743 (0.058)	7.212 (0.356)

Bootstrapped p-value in parenthesis for 500 replications

Test results for model 1 clearly reject the linear model to the benefit of a non-linear model of adjustment with two thresholds. Results are similar when introducing the consumer price index into the cointegrating equation (model 2) although somewhat ambiguous. The linear model is rejected but the number of regimes in the alternative is not clearly identified by the tests.

The TAR(3) model estimates for the residual of the long run relationship are given in table 10; corresponding ECM estimates are in table 11. Estimation results confirm the previous ones.  $\mu_t$  is not stationary and cointegration is inactive as long as discrepancies from long term equilibrium lie within the band defined by the two thresholds; the process is mean reverting in the outer regimes. We note that the speed of adjustment in the lower regime is not significantly higher than in the upper regime (Table 11). This may be the consequence of the low number of observations in each regime that can affect the precision of parameter estimate.

Table 10. Estimate results from TAR(3;1,1). Dependent variable:  $\Delta\mu_t$

		regime			Wald $\rho^l=\rho^m=\rho^u=0$ F-stat	Wald $\rho^l=\rho^u$ F-stat	$\tau_1$	$\tau_2$
		Lower	Middle	Upper				
Model 1	$\rho$	-0.223	0.352	-0.097	7.096	1.30	-27.726	32.823
		(-2.161)	(2.992)	(-2.532)	[0.000]	[0.259]		
		[0.035]	[0.004]	[0.014]				
	No obs	12	31	18				
Model 2	$\rho$	-0.403	0.345	-0.326	9.42	0.163	-0.094	0.023
		(-2.547)	(1.513)	(-3.814)	[0.000]	[0.688]		
		[0.014]	[0.136]	[0.000]				
	No obs	9	31	15				

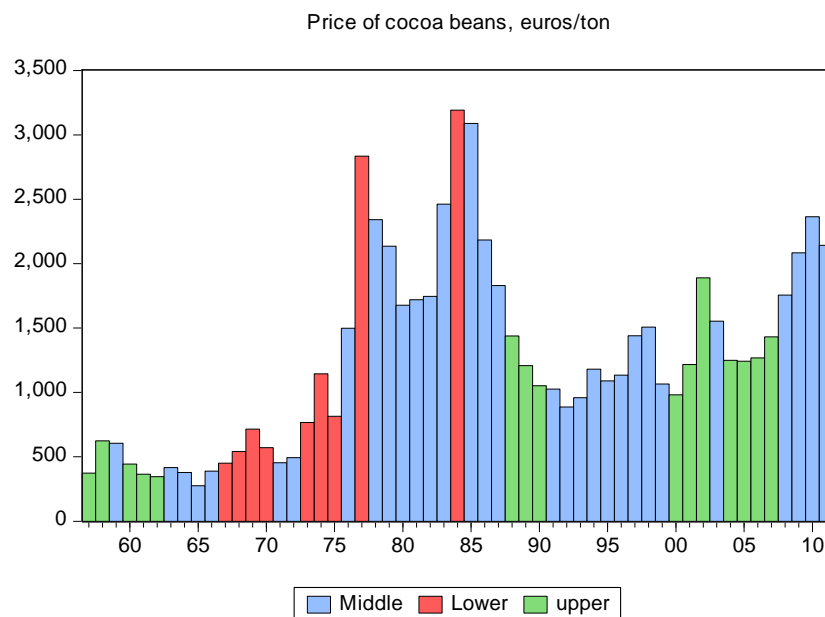
Number of lags: 1.  $t$ -stats are in parenthesis; p-value are in brackets

Table 11. Estimate results for the 3 regimes Error Correction Model

Dependent variable: $\Delta P_c$		Regime			Wald tests ( $F$ -stat)		
	lower	middle	upper	Adj R <sup>2</sup>	$\phi^l = \phi^m = \phi^u = 0$	$\phi^l = \phi^m = \phi^u$	$\phi^l = \phi^u$
	$\phi^l$	$\phi^m$	$\phi^u$				
Model 1	-0.068	0.026	0.006	0.288	5.35	6.64	8.39
(P-value)	(0.001)	(0.332)	(0.656)		(0.003)	(0.003)	(0.005)
Model 2	-0.529	0.267	-0.441	0.608	10.89	5.98	0.16
(P-value)	(0.001)	(0.288)	(0.004)		(0.000)	(0.005)	(0.687)

no lag in first difference variables.

Figure 4. Timing of regime switching (annual data, Model 2)



The broad picture that emerges when looking at the timing of the three regimes (figure 4) is the same than previously observed with monthly data (figure 3). Results clearly show that the historical booms in the cocoa price recorded in 1974, 1977 and 1984 were transmitted to the consumer (red periods). Between the 1977 and 1984 booms the price disequilibrium remained at a relatively low level. Prices didn't adjust despite the 1980 fall in the cocoa price. The phase of sharp and prolonged decline in world prices that started in 1986 resulted in large price disequilibrium – retail price being above their long run mean - that was corrected only at the end of the 80s. Results also confirm that over the entire 1990-2000 period the chocolate consumers have been fully isolated from cocoa price fluctuations. During this period, the price disequilibrium was not large enough to trigger an adjustment of the retail price. The price of the chocolate bar did not adjust to the 1999-00 fall in cocoa price. Correction



occurred at the beginning of the 2000s. However the chocolate retail price was still above its long run value in 2007 so that the 2010 cocoa price recovery was not passed on to consumers.

## **5. Concluding remarks**

The analysis of the relationship between the cocoa price and the chocolate price, in the long and short run, clearly leads to rejecting the hypothesis of linear price adjustment and highlights two types of inefficiency in the cocoa-chocolate chain.

First, the retail consumer price does not adjust to small positive or negative shocks in the price of cocoa beans. This phenomenon can be interpreted as reflecting the presence of menu costs. Menu costs create a “band of inaction” within which price disequilibria are not corrected. On the 1955 -2011 period, the no cointegration regime is the dominant one signaling important, may be excessive, adjustment costs.

Secondly, the retail price adjusts faster to large increases in the price of the raw material than to large decreases. The consequence is that the chocolate price corrects quickly a disequilibrium following a boom in the cocoa price but reverts back more slowly to its long term value when the cocoa price declines.

The importance of adjustment costs and the asymmetric transmission of positive and negative price shocks to consumers can be seen as evidence of inefficiencies in the processing, manufacturing and distribution of chocolate products. It can also be the manifestation of non-competitive behavior in the chocolate industry which is highly concentrated.

Another interpretation is that retail chocolate price rigidity reflects the price smoothing strategy of chocolate manufacturers. Inventory management may allow firms to dampen cocoa price shocks and protect risk adverse consumers against fluctuations in the price of the raw material. However, retail prices are more rigid downward than upward for input shocks meaning that consumers bear the burden of an asymmetric price smoothing strategy. The analysis has shown that the retail price of the chocolate bar adjusted rapidly to large cocoa increases during the first half of the period under study but tended to remain above its long run mean during the long phase of low cocoa prices that began in 1985.

One of the limits of the analysis is that the adjustment costs are supposed to be constant over the whole period. The methodology does not allow to catch a potential shift in agents' behavior in line with the increasing concentration of the chocolate industry. Taking into account for such a break would be difficult due to the relatively low number of observations.

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